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Contents

Abstract

The study investigates long-run relationships between futures and spot prices of cocoa on the New York CSCE and London Fox, respectively, and between both markets. By means of the Johansen Maximum Likelihood approach and the inclusion of interest rates as conditioning variables, the three hypothesized cointegrating vectors are obtained. It turns out that the usage of interest rates is crucial for detecting long-run stationary relationships between spot and futures prices on individual markets. This might explain the failure of previous studies to discover cointegration between spot and futures prices on commodity markets. The existence of asymmetries in the response to deviations from equilibrium relationships is also observed: Futures prices Granger-cause spot prices, but not vice versa. This result is interpreted as evidence for spot prices to react slowly to new information.

Zusammenfassung

Diese Studie untersucht die langfristigen Beziehungen von Futures- zu Spot-Preisen für Kakao an der New Yorker CSCE und der London Fox, sowie zwischen diesen beiden Märkten. Mittels der Maximum-Likelihood-Methode nach Johansen und unter Einbeziehung von Zinsen als konditionierende Variablen können die drei erwarteten Kointegrationsvektoren ermittelt werden. Es stellt sich heraus, daß die Verwendung von Zinsen für die Auffindung von langfristigen stationären Beziehungen zwischen Spot- und Futures-Preisen auf einzelnen Märkten von äußerster Bedeutung ist. Das könnte das Versagen früherer Studien erklären, die die Untersuchung von Kointegration zwischen Spot- und Futures-Preisen auf Rohstoffmärkten zum Ziel hatten. Die Existenz von Asymmetrien als Reaktion von Abweichungen von Gleichgewichtsrelationen kann ebenfalls beobachtet werden: Furtures-Preise sind Granger-kausal für Spot-Preise, aber nicht umgekehrt. Dieses Ergebnis kann als Hinweis dafür gewertet werden, daß Spot-Preise langsam auf neue Informationen reagieren.

JEL-Classification: C32, G15

1 Introduction

Tests of the efficiency of arbitrage in commodity markets generally provide mixed results. The tests have either been based on the law of one price (LOP) or have examined relationships between spot and futures prices. The LOP states that the price of a homogeneous product (expressed in a single currency) must be the same in all markets after transaction costs have been accounted for. Futures arbitrage implies that futures prices must equal expected future spot prices corrected for the opportunity costs of holding the commodity.

Several recent studies have taken a new look at this issue using cointegration techniques (e.g., Ardeni, 1989; Baffes, 1991; Baillie, 1989; Baillie and Myers, 1991; and Karbuz and Jumah, 1995; among others). The rationale behind this approach is the widely acknowledged fact that commodity prices follow nonstationary processes. From this point of view, arbitrage between several assets implies that their prices move together, such that the price differences are represented by a stationary process, i.e., the prices are cointegrated. Though this proposition is a rather weak implication of arbitrage, various studies found mixed evidence for the existence of cointegration relationships. Ardeni (1989) examined arbitrage between different markets using import prices and found cointegration for only three out of fifteen commodities. Baillie and Myers (1991) found no cointegration between spot and futures prices for six commodities. Using cocoa and coffee prices on the New York Coffee, Sugar and Cocoa Exchange (New York CSCE) and London Futures and Options Exchange (London Fox) Karbuz and Jumah (1995) found cointegration between markets, thus supporting the LOP, but weak support for stationary long-run relationships between futures and spot prices.

There are two possible reasons for the failure to find cointegration in previous studies. The first concerns the neglect of interest rates which, as they enter the arbitrage relationships between spot and futures prices, clearly play an important role. As interest rates themselves follow nonstationary processes their omission may severely affect the results of corresponding cointegration tests. Barnett, Bessler and Thompson (1983) – among others – have also shown that commodity prices are sensitive to monetary factors. The second reason is associated with the reliance on single-equation techniques for cointegration analysis. The aforementioned studies mostly used the method proposed by Engle and Granger (1987) to examine pairwise cointegration between the series in question. The Engle-Granger technique does not provide fully efficient estimates (see e.g., Banerjee et al., 1993). In particular, Hendry and Mizon (1992) have emphasized that single equation estimation techniques are inefficient if there are other series left out that are themselves cointegrated with the series in question. In the present context this regards the mutual relationships between spot and futures prices on individual markets. If, for instance, there exists cointegration between individual markets, then a single equation cointegration test between spot and futures prices on only one market might be an inefficient tool. This calls therefore for the application of a full systems approach.

Using monthly spot and futures prices on the New York CSCE and the London Fox we test simultaneously for both the LOP and futures arbitrage. To account for interest rates, 3-month US and UK interest rates are added as conditioning variables to the system, following the approach of Johansen (1992a). The resulting estimates are compared with a model without interest rates and the Engle-Granger approach. Finally we conduct various tests of structural hypotheses on the cointegration relationships and the adjustment of spot and futures prices to deviations from long-run equilibrium.

2 Arbitrage in Commodity Markets

Economic theory postulates that the efficient functioning of asset markets should prevent incidences of arbitrage profit opportunities from occurring. The first type of arbitrage – futures arbitrage – implies a certain equilibrium relationship between futures and spot prices. In case of a deviation from the equilibrium an investor could gain risk-less profits by taking opposite long and short positions in spot and futures trading. The relationship between futures and spot prices can be demonstrated by the arbitrage pricing model by Hull (1989). Let s_t and $f_{t,k}$ be the respective spot and futures prices (in logs) of a security at time t and for delivery at time $t+k$. $r_{t,k}$ denotes the return on a riskless k-period bond issued at time t. For a security that provides no income the indifference between holding the security or a futures contract, corrected for the loss in interest income when the security is held, implies that in equilibrium $f_{t,k} = s_t - r_{t,k}$. In the case of commodity markets the arbitrage relationship has to be extended in two respects: First, it has to include storage costs SC that arise when the commodity is held. Second, for commodities that are, to a significant extent, held for consumption the so-called convenience yield CY must be taken into account. Firms that keep the commodity in inventory do so because of its consumption value, i.e., its value for the production of final goods. They are reluctant to sell the commodity and buy futures contracts since futures contracts cannot be consumed. Thus a firm may not fully take advantage of arbitrage opportunities that arise when the futures price is below the equilibrium relationship. These considerations lead to the arbitrage relationship (see Hull, 1989, p. 48ff)

$$f_{t,k} = s_t - r_{t,k} + \mu_k, \quad \mu_k = SC - CY. \tag{1}$$

Note that, since arbitrage relies purely on risk-neutral portfolio adjustment, equation (1) does not depend on the risk neutrality of the investors. In a different view, not involving arbitrage arguments, futures markets are seen as insurance markets allowing diversification of spot price risk. Investors take a decision between buying at time t a futures contract that expires at $t+k$ or buying the commodity at time $t+k$. In this case the equilibrium relationship between futures and expected future spot prices is stated as

$$f_{t,k} = E_t s_{t+k} - r_{t,k} + \pi_k, \tag{2}$$

where $E_t s_{t+k}$ is the expectation at time t of the spot price at $t+k$. π_k denotes the risk premium which a risk-averse investor demands for selling a futures contract thereby providing insurance against unexpected changes in spot prices (see e.g., Kaminsky and Kumar, 1990). Note that in the case of risk-neutral investors $\pi_k = 0$. Note also that in the case of the efficiency of the spot market itself, $E_t s_{t+k} = s_t$ (see e.g., LeRoy, 1989).

The second type of arbitrage in commodity markets concerns spot prices between markets. The LOP implies that international commodity arbitrage in efficient markets leads to an equalization of commodity prices across different markets and can be represented as:

$$s_t = s_t^* + e_t + \theta \tag{3}$$

where s_t and s_t^* are the commodity prices in countries 1 and 2 (both in domestic currency) and e_t is the spot exchange rate of the two currencies (variables are in logs). The constant θ represents transaction costs like transport costs and shifting effects (import quotas, levies, etc). The LOP forms the basis of Purchasing Power Parity theory and the determination of exchange rates.

Futures arbitrage and the LOP between spot prices also imply the existence of equilibrium relationships between futures prices on different markets when exchange rate expectations are taken into account. This can be seen from the following set of equation s:

$$\begin{array}{lclclcl} f_{t,k} & = & E_t s_{t+k} & - & r_{t,k} & + & \pi_k \\ f_{t,k}^* & = & E_t s_{t+k}^* & - & r_{t,k}^* & + & \pi_k^* \\ s_t & = & s_t^* & + & e_t & + & \theta \end{array} \tag{4}$$

which implies

$$f_{t,k} - f_{t,k}^* = E_t(s_{t+k} - s_{t+k}^*) - (r_{t,k} - r_{t,k}^*) - (\pi_k - \pi_k^*) = E_t(s_{t+k} - s_{t+k}^* - \hat{e}_{t+k}) - (\pi_k - \pi_k^*),$$

where the relationship $E_t \hat{e}_{t+k} = r_{t,k} - r_{t,k}^*$ is the uncovered interest parity condition.

There is a significant amount of literature in testing equations (1) to (3). The existence of a risk premium and the question of whether futures prices are unbiased predictors of future spot prices (and also future futures prices) as implied by equation (2) for $\pi_k = 0$ have been investigated e.g., by Hazuka (1984), Jagannathan (1985) and Kaminsky and Kumar (1990). The latter conclude that the evidence for the existence of a risk premium is mixed and that it might vary over time. Inefficiencies in commodity arbitrage between markets have been linked to the presence of trade barriers, explicit and implicit contracts (Ardeni, 1989), imperfect competition in goods markets (Stigler and Sherwin, 1985) and product differentiation (Karbuz and Jumah, 1995).

It is important to note that cointegration is a rather weak consequence of market efficiency. Market efficiency means by definition that the equilibrium errors in equations (1) to (3) are unforecastable. Cointegration only implies that the equilibrium errors follow a stationary process. Thus, the existence of a cointegration relationship can only be interpreted as an equilibrium relationship that holds in the long run, whereas there might be significantly forecastable deviations in the short run.

3 Methodology

The cointegration analysis is based on the approach initiated by Johansen (1988). Consider an n-dimensional stochastic process x_t which is generated by a Gaussian (k+1)-th order vector-autoregression:

$$A(L)x_t = \mu + \epsilon_t, \tag{5}$$

where $x_t = (x_{1,t}, \ldots, x_{n,t})'$ and $A(L) = I_n - \sum_{m=1}^{k+1} A_m L^m$. ϵ_t is normally distributed with $N(0, \Lambda)$. It is assumed that the roots of the characteristic polynomial are strictly outside the unit circle or equal to one, i.e., $|A(L)| \neq 0$ for $|z| < 1$. If A(L) contains no unit root, i. e., $|A(L)| \neq 0$ for $|z| = 1$, the VAR in (5) is stable and the process X_t is stationary. In particular, X_t fluctuates around a constant mean and the impact of innovations ϵ_t on X_t is zero in the long run.

Generally, the polynomial A(L) may contain $0 \leq s \leq n$ unit roots.[1] This is equivalent to the existence of $p = n - s$ stationary independent linear combinations $\beta_i' X_t$, i.e., cointegration relationships and s independent nonstationary stochastic trends. More specifically, Engle and Granger (1987) have shown that the process (1) has an error-correction (EC) representation of the form

$$A^*(L)\Delta x_t = \mu + \alpha\beta' x_{t-k-1} + \epsilon_t = \mu + \Pi x_{t-k-1} + \epsilon_t, \tag{6}$$

where the rank of the n x n matrix Π equals the number of cointegrating vectors $p <= n$ in the system. Thus Π can be written as $\Pi = \alpha\beta'$, where α and β are both of the dimension n x p and rank p. The matrix β contains the cointegrating vectors, i.e., $\beta = (\beta_1, \ldots \beta_p)$. The matrix of adjustment coefficients α describes the speed of adjustment of the particular series $x_{i,t}$ to deviations from the cointegration relationships, i.e ., the equilibrium errors.

Engle and Granger (1987) proposed a two-step estimator for equation (6). The first step involves a linear regression of $x_{1,t}$ on $x_t^{(1)} = (x_{2,t}, \ldots, x_{n_t})'$ and a subsequent test for cointegration. In the case of cointegration the estimated equilibrium error $v_{t-k-1} = x_{1,t-k-1} - \hat{\beta}' x_{t-k-1}^{(1)}$ is substituted for $\beta' x_{t-k-1}$ in equation (6) and α and $A^*(L)$ are estimated by Ordinary Least Squares. While, for the case of $n > 2$, it is possible to test for the number p of cointegrating relationships by a sequence of recursive Engle-Granger-tests, this is a rather inefficient approach (see e.g., Phillips, 1991).

For the general multivariate case Johansen (1988) derived a maximum likelihood estimation technique based on reduced rank regression techniques and a trace test for the rank of Π. The ML-procedure determines the cointegration space for an n-dimensional system simultaneously and thus provides a complete assessment of the number of cointegration relationships. It is important to note, however, that though the analysis yields the space spanned by the cointegrating vectors, the particular

[1]It is also assumed that X_t is first difference stationary which implies that $A^*(L)$ of equation (6) contains no more unit roots.

cointegrating vectors are not uniquely identified. More specifically, α and β are subject to rotation, i.e., $\Pi = \alpha\beta' = \alpha\xi\xi^{-1}\beta'$ for any invertible p x p matrix ξ.

In this study the Johansen ML-procedure is applied to test simultaneously for futures arbitrage on the New York CSCE and London Fox and for the LOP for spot and futures prices between both commodity exhanges. The LOP and futures arbitrage together imply the existence of three independent cointegrating vectors, describing the stationary relationships between New York and London prices as well as futures and spot prices on both markets. We estimate the system of the four variables:

- futures prices New York (FNY)
- futures prices London (FL)
- spot prices New York (SNY)
- spot prices London (SL)

Furthermore, we add 3-month US and UK treasury bill rates as exogenous variables to the cointegration relationships. The neglect of interest rates might be responsible for the failure to find the hypothesized cointegration relationships in previous studies. As long as interest rates follow a stationary process, their omission does not severely affect the outcome of cointegration tests. Specifically, the estimates of β remain consistent if the short run is misspecified. If interest rates follow a nonstationary process, however, which has been shown to be the case by numerous studies (see e.g., Engle, Lilien and Robins, 1987) their omission leads to a misspecification of the cointegrating vectors themselves. This has already been pointed out by Baillie and Myers (1991), who, however, left interest rates out in their estimations. The inclusion of interest rates leads to a six-dimensional system of the form

$$\begin{pmatrix} A_{11}^*(L) & A_{12}^*(L) \\ A_{21}^*(L) & A_{22}^*(L) \end{pmatrix} \begin{pmatrix} \Delta y_t \\ \Delta r_t \end{pmatrix} = \mu + \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{pmatrix} \beta' \begin{pmatrix} y_{t-k-1} \\ r_{t-k-1} \end{pmatrix} + \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix} \quad (7)$$

with the 4 x 1 vector y_t of the various cocoa prices and the 2 x 1 vector r_t of US and UK interest rates. However, as pointed out by Johansen (1992a), the number of parameters explodes in large systems with a corresponding loss of efficiency in subsequent cointegration tests. It is thus preferable to estimate the submodel for y_t conditional on interest rates r_t

$$\tilde{A}_{11}^{*}(L)\Delta y_t = \tilde{\mu} + \tilde{A}_0\Delta r_t + \tilde{A}_{12,l}^{*}(L)\Delta r_{t-1} + (\tilde{\alpha_{11}}\,\tilde{\alpha_{12}})\,\beta' x_{t-k-1} + \tilde{\epsilon}_{1,t} \tag{8}$$

where the vector x_t contains both $x_t = (y_t', r_t')'$. Equation (8) is devoted to the conditional systems approach advocated by Johansen (1992a) and Hendry and Mizon (1992). Johansen (1992a) has shown that, if in equation (7) both $\alpha_{21} = 0$ and $\alpha_{22} = 0$, that is, if Δr_t does not respond to deviations from long-run equilibrium $\beta' x_{t-k-1}$, then the ML-estimates of α and β are identical for both models (7) and (8). In such a situation the vector r_t is called weakly exogenous with respect to α and β (Engle, Hendry and Richard, 1983). If the above condition is not satisfied, then the estimation of the partial model is an inefficient tool for the determination of the cointegration rank. If in equation (7) $\alpha_{21} = \alpha_{22} = A_{21}^{*}(L) = A_{22}^{*}(L) = 0$ then r_t is weakly exogenous with respect to all parameters in equation (5) (see e.g., Mosconi and Giannini, 1992). In this case there is no loss of information in estimating the conditional model (8). In large systems where some variables are weakly exogenous, the estimation of a conditional model might bring considerable efficiency gains as compared to model (7) (see e.g., Urbain, 1993). In case some variables are likely to be weakly exogenous, Johansen (1992a) recommends the estimation of the conditional model (8) and a subsequent misspecification test for the weak exogeneity. Also, as pointed out by Urbain (1993), the estimation of the conditional model (8) has the further advantage that it does not rely on the correct specification of the process of the exogenous variables. Dengsoe et al. (1994) provide critical values for the trace statistics of the conditional model.

The above discussion also points to a further weakness of existing work on cointegration in commodity markets as concerns the validity of single equation analysis. This may be seen, e.g., in the case of testing for the LOP by Karbuz and Jumah (1995), as the estimation of a partial system of two spot prices s_t in the context of a full system $x_t = (s_t, f_t)'$ that also includes futures prices. As long as futures prices are not weakly exogenous to spot prices, bivariate cointegration tests based on the partial system are inefficient. Equivalently, a test for cointegration between futures and spot prices on one single market might be inefficient as long as they also react to equilibrium errors in cointegration relationships with other markets.

4 Results

4.1 Cointegration Analysis

We use monthly average spot and futures prices of cocoa for the period of 1981:1 to 1991:12 between the New York CSCE and London Fox which are worldwide the two most important trading centers for cocoa. The prices are logarithmized.[2] Augmented Dickey-Fuller (ADF) tests for the existence of unit roots in the series (Dickey and Fuller, 1979, 1981) are shown in Table 1. The tests indicate the presence of a unit root for each of the commodity prices as well as interest rates. Stationarity is attained for all series after first differencing which implies that all variables are integrated of order one. Table 2 shows the results of Engle-Granger tests for pairwise cointegration. We find cointegration between markets for both spot prices and futures prices at the 5% level, but no evidence for cointegration between spot and futures prices on the respective commodity exchanges.[3] This also holds if interest rates are included in the cointegration equations.

We now turn to full-systems analysis using the Johansen ML-procedure. Boswijk and Franses (1992) have found that overfitting in a vector error-correction model implies a loss in power for subsequent cointegration tests, while underfitting leads to spurious cointegration. For the determination of lag length in (6) they thus suggest the application of the Johansen ML-procedure for different lag lengths and

[2]Source of the cocoa prices is the ICCO-statistics. 3-month treasury bill rates are from the International Financial Statistics (IMF).

[3]Using the same data set with a somewhat longer estimation period Karbuz and Jumah (1995) found cointegration between spots and futures at the 10% level.

to base the final choice on the absence of serial correlation in the residuals and the significance of parameters at higher lags using LR-tests. In the case of changes in cocoa prices under the hypothesis of efficient markets one would expect a rather low lag length. It turns out, however, that higher lags are needed to remove residual serial correlation. The application of the above procedure results in a model with four lags for cocoa prices and two lags for interest rates. The analysis also showed two significant positive outliers for London futures prices at the beginning of the sample (81:3 and 81:7) which lead to a rejection of normality of the residuals in the corresponding equation (compare Tables 3 and 4). A dummy is therefore included as a proxy for these two data points.

The basic results for the full systems approach are presented in Tables 3 and 4 for both model (6) and the conditional model (8), respectively. Dengsoe et al. (1994) provide critical values for the trace statistics of the conditional model only for the case where there is no deterministic trend in the data so that the constant fully enters the cointegration relationships, i.e., $\mu = \tilde{\alpha}_0\beta$ (see e.g., Johansen, 1991). Consequently, we apply a test of that restriction. The first line of both tables shows the LR-statistics of the hypothesis $\mu = \tilde{\alpha}_0\beta$, which is χ^2-distributed with four degrees of freedom under the null. The restriction is rejected for model (6), but is accepted at a reasonable significance level for model (8).

Tables 3 and 4 contain the singular values λ_i of Π and the trace statistics $\sum_{j=1}^{i} - \ln(1-\lambda_j)$, which is used to test for the dimension of the cointegration space, together with its critical values. From the analysis, model (8) yields three cointegrating vectors at the 5% level. In contrast, for model (6) we obtain only one cointegrating vector significant at the 5%-level and a second at the 10%-level.[4] It is concluded that the inclusion of interest rates considerably improves the estimates of the long-run relationships between cocoa prices and is necessary for establishing cointegration between spot and futures prices on commodity markets.

[4]The results for both models remain essentially unchanged for higher lag lengths. It is also noticeable that the efficiency loss in the estimation of the full 6-dimensional system (7) is that large so that only one cointegrating vector is found.

For the misspecification test of weak exogeneity of interest rates with respect to α and β, interest rates (in first differences) are regressed on four autoregressive lags and the estimates of the cointegrating vectors $\hat{\beta}' x_{t-k-1}$ (see Johansen, 1992a). The Wald test yields low F-statistics for the significance of the latter.[5] For the test of weak exogeneity with respect to all parameters in model (7) lagged cocoa prices (in first differences) are added to the above equations. Again, there is no evidence for an influence of cocoa prices on interest rates. Thus, the ML-estimates of all parameter values of the conditional model can be considered as being unbiased with respect to the full model (7) for (y_t, r_t).[6]

4.2 Structural Tests on the Cointegrating Vectors and Adjustment Coefficients

We proceed with tests on various overidentifying restrictions of β for the conditional model and now restrict the constant to enter the cointegration relationships, i.e., $\mu = \tilde{\alpha}_0 \beta$. Johansen (1992b) has proposed a switching algorithm for estimating equation (8) under certain restrictions on the particular cointegrating vectors, i.e., the combined hypotheses

$$\beta = (\beta_1, \beta_2, \ldots, \beta_p) = (H_1\phi_1, H_2\phi_2, \ldots, H_p\phi_p) \tag{9}$$

with fixed restricting n x q_i matrices H_i and the q_i x 1 vectors of freely estimated parameters ϕ_i for each cointegration relationship. The LR-statistics for the combined hypothesis $H_i (i = 1, \ldots, p)$, given the existence of three cointegrating vectors, is asymptotically standard χ^2_k-distributed, where k denotes the total number of restrictions on β.

Table 5 presents estimates of β using six just-identifying zero restrictions (H_0^β) on cocoa price coefficients that give a representation of the cointegrating vectors

[5]The values of the F(120,3)-statistics are 0.96 (p = .41) and 0.65 (p =.58) for US and UK interest rates, respectively.

[6]There is also no evidence for a cointegration relationship between r_{US} and r_{UK}.

according to equations (4). The first cointegrating vector β_S represents a relationship between spot prices, the second and third (β_{NY} and β_L) describe the relationships between futures and spot prices on either market. In sum, the estimates are quite close to the theoretical relationships (4). The freely estimated cocoa price coefficients are close to unity in absolute value. The interest rate coefficients are relatively small for the relationship between spot prices β_S. With respect to the relationship between futures and spot prices, UK interest rates were found to have a (wrong) negative sign on both markets.[7]

Initially, we test for the coefficients of cocoa prices being all equal to either one or zero, as implied by equations (4). The coefficients for interest rates and the constant are left unrestricted. Note that the outcomes of this and the subsequent tests are independent of the specific structure imposed on β as long as the restrictions are accordingly specified. Though the unrestricted estimates are close to unity, this hypothesis (H_1^β in Table 6) is very close to rejection at the 5%-level. The restrictions H_c^β and H_r^β of zero coefficients for the constant and interest rates, respectively, are tested separately for β_S and both futures-spot relationships β_{NY} and β_L. Both restrictions ($H_{c,23}^\beta$ and $H_{r,23}^\beta$) are rejected for β_{NY} and β_L, while they are accepted for β_S ($H_{c,1}^\beta$ and $H_{r,1}^\beta$). Finally, Table 8 presents the estimates for β_{rc} under the full set of restrictions H_{rc}^β implied by equations (4), where cocoa prices are not restricted to unity. The restrictions are accepted at a fairly high significance level. Both interest rate coefficients have the correct sign.

Following the approach proposed by Johansen and Juselius (1992) we proceed by testing various structural hypothesis on the adjustment coefficients $\tilde{\alpha}$ based on fixed estimates of the just-identified cointegrating vectors according to H_0^β. The results of the corresponding Wald tests are presented in Table 7.[8] The first test examines the null hypothesis H_i^α of particular rows of $\tilde{\alpha}$ being zero, which means that the corresponding variable does not respond to any of the three equilibrium errors.

[7]Note that the coefficients for 3-month interest rates should be equal to one only for a futures with a maturity of 3 months.

[8]Johansen and Juselius (1990) have shown that the standard asymptotic distribution results apply for tests on α. Sims' (1980) small sample correction was used.

The null hypothesis is rejected for both spot (at the 10%-level) but not for futures prices. However, it cannot be true for both futures prices.[9] We therefore test two sub-hypotheses for each row of $\tilde{\alpha}$. For the first ($H^{\alpha}_{i,1}$) the adjustment coefficients for β_S are set to zero. This restriction is close to rejection for both New York prices, but is not rejected for London prices. The second test ($H^{\alpha}_{i,23}$) puts zero restrictions on the adjustment coefficients for both β_{NY} and β_L. These restrictions are rejected for both spot prices but not for futures prices. In sum, the results suggest the existence of two asymmetries in the adjustment to long-run equilibrium: Futures prices do not respond to equilibrium errors in the futures-spot price relationships and London prices do not respond to equilibrium errors between markets β_S.

These tests are not completely conclusive, however, as concerns the interpretation of mutual dependencies between the particular prices. The non-reaction of London prices to β_S is not equivalent to the proposition that London prices do not respond to New York prices, since London prices might respond to β_{NY}. Similarly, since futures prices respond to β_S – based on the above tests – it is not possible to conclude that futures prices do not respond to spot prices. Moreover there might be dependencies due to the lagged cocoa prices in first differences, i.e., the lag polynomial $\tilde{A}^*_{11}(L)$ in equation (7). A more rigorous test of a set P_1 of prices not responding to a set P_2 can be performed by the variant of Granger-causality tests proposed by Toda and Phillips (1993). Toda and Phillips (1993) have pointed out that in an error-correction-model (6) the proposition of y_j not Granger-causing y_i is equivalent to both $A^*_{ij}(L) = \Pi_{ij} = 0$ and proposed to test each subhypothesis sequentially. Also, they derived rank conditions for the Wald statistics to be standard χ^2-distributed.

We test for the combined hypotheses of mutual Granger-causality between spot and futures prices and Granger-causality between New York and London, respectively. While the various restrictions ($H^{a}_{(.)}$) on the lag polynomial $\tilde{A}^*_{11}$ in equation (8) are straightforward the restrictions ($H^{\alpha}_{(.)}$) on $\tilde{\Pi} = \tilde{\alpha}\beta$ need some consideration. On the basis of the just-identified representation of β shown in Table 5, however, the zero restrictions on $\tilde{\Pi}$ can be readily transformed into equivalent restrictions on

[9] This would imply that the rank $rk(\tilde{\Pi}) = rk(\tilde{\alpha}\beta) \leq 2$.

$\tilde{\alpha}$. For the test of New York not Granger-causing London (H_L^α), for instance, we put zero restrictions on both London price equations for the adjustment coefficients corresponding to β_S and β_{NY}, while the adjustment coefficients corresponding to β_L are left unrestricted. For the test of futures prices not Granger-causing spot prices (H_S^α), adjustment coefficients for β_{NY} and β_L are set to zero for both spot price equations (see Table 7). For the reverse test of spot prices not Granger-causing futures prices (H_F^α), we use a different representation of β where β_S is substituted by the cointegration relationship between futures prices β_F.[10]. It can also be shown that the rank conditions for the feasibility of Granger-causality tests given by Toda and Phillips (1993) are satisfied.

The results confirm the existence of the above mentioned asymmetries, in particular as concerns the dependencies between spot and futures prices. Futures prices Granger-cause spot prices: The restrictions imposed on $\tilde{\alpha}$ and $\tilde{A}^*(L)$ are both rejected at the 5 %-level. The reverse hypothesis of spot prices not Granger-causing futures prices cannot be rejected. The asymmetries between the London Fox and the New York CSCE are confirmed only for the adjustment coefficients, but not for the lag polynomials. The hypothesis H_{NY}^α that the New York CSCE neither responds to equilibrium errors between markets (β_S) nor to equilibrium errors between London spot and futures prices (β_L) is rejected at the 1%-level. The reverse hypothesis H_L^α for the London Fox is not rejected. However, both corresponding restrictions on the lag polynomials (H_{NY}^a and H_L^a) are rejected at the 1 %-level. Thus the tests indicate mutual Granger-causality between London and Ne w York. Table 7 shows also the Wald tests for the combined restrictions $H_{NY;F}^\alpha$ and $H_{NY;F}^a$, while Table 8 presents the restricted estimates of $\tilde{\alpha}$ for β according to H_{rc}^β, where again β_S is replaced by β_F (no restrictions on the lag polynomials are imposed). The graphs of the cointegrating relations are shown in Fig. 2.

[10]This is found by an estimation using different just-identifying restrictions and is equal to $(1, -1.023, 0, 0, -0.0004, 0.037, 0.048)$.

5 Conclusions

The full system analysis of New York and London spot and futures prices for cocoa confirms the existence of the hypothesized three independent cointegration relationships. The study concludes that the LOP for cocoa holds in the long run. Interest rates were found to play a key role in establishing stationary long-run relationships between spot and futures prices of cocoa on the individual markets. The neglect of interest rates could be the reason why previous studies failed to detect stationary long-run relationships between spot and futures prices for different commodities. The results also proved that the Johansen ML-estimation is superior to the single-equation Engle-Granger-approach.

Tests on the coefficients of the estimated long-run relationships confirm the structure imposed by the arbitrage relationships (4) to a large extent. The cocoa price coefficients are close to unity while the interest rate coefficients have the correct sign. There is, however, some ambiguity concerning the sign of the constant (the risk premium) in the futures-spot price relationships, which is positive for New York but is negative for London. This difference in risk premia between markets corresponds to the results obtained by Kaminsky and Kumar (1990) who found risk-premia to vary over time becoming negative for certain periods.

The study also revealed evidence of two asymmetries in the markets as concerns the adjustment to long-run equilibrium. First, deviations from long-run equilibrium between markets are offset predominantly by an adjustment of New York prices. Second, deviations from long-run equilibrium in future-spot price relationships are offset predominantly by an adjustment of spot prices. The asymmetry between the New York and London markets is in accordance with the larger trading volume of the London Fox. However, the above result is weakened by the existence of Granger-causality in both directions. This implies that both markets influence each other in the short run.

On the other hand, the asymmetry between spots and futures prices holds both in the short and the long run. One possible explanation of the latter result is that futures prices adjust more quickly to new information than spot prices. Speculation in commodity markets is concentrated in futures. Futures prices might quickly reflect new information on future demand and supply conditions, for instance due to current weather conditions, and therefore are equal to expectations on future spot prices, as implied by the efficiency hypothesis, while intertemporal arbitrage may hold for spot prices only to a limited extent. There are several possible reasons for both inelastic supply of and demand for cocoa. Inelastic supply may stem from the fact that suppliers want to smooth their income stream. In many of the supplier countries, the production of cocoa is the backbone of the domestic economy. Cocoa exports are quite often the most important tax base of these countries. Income changes can thus directly influence government budgets. Thus, in case of an expected increase in spot prices, the supplier countries are likely to smooth their income stream by selling below the expected future spot price. Inelastic demand may be due to storage costs and low price and income inelasticities of demand for final cocoa products. For more detailed discussions of inelastic demand and supply on cocoa markets see e.g., Akiyama and Duncan (1981) and Jumah (1986).

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TABLE 1: AUGMENTED DICKEY FULLER – TESTS

	Levels t-Value	Levels (Lags)	First Differences t-Value	First Differences (Lags)
FNY	-1.43	(4)	-7.99	(2)
FL	-1.18	(3)	-7.91	(2)
SNY	-1.21	(3)	-7.82	(2)
SL	-1.28	(1)	-9.82	(0)
r_{US}	-2.07	(0)	-10.34	(0)
r_{UK}	-2.28	(0)	-11.08	(0)

Note: 5% and 10% critical (absolute) values are 2.86 and 2.56, respectively (McKinnon, 1991).

TABLE 2: PAIRWISE ENGLE-GRANGER-COINTEGRATION TESTS

			Long-Run Coefficients		t-Value	(Lags)
FNY,	FL		1.041		-4.34	(0)
SNY,	SL		0.978		-4.74	(2)
FNY,	SNY		0.994		-2.77	(0)
FL,	SL		0.934		-2.83	(2)
FNY,	SNY,	r_{US}	1.011	-0.0012	2.87	(0)
FL,	SL,	r_{UK}	0.912	-0.0017	2.83	(2)

Note: 5% and 10% critical (absolute) values are 3.35 and 3.04, respectively (McKinnon, 1991).

TABLE 3: ML-COINTEGRATION ANALYSIS OF MODEL (5)

LR-test of $\mu = \tilde{\alpha}_0\beta$ (df = 1)			11.22
LR-test of dummy (df = 4)			28.12
	Eigenvalues		
0.203	0.136	0.058	0.013
	Trace Statistics		Critical Values
	1.71	$r \leq 3$	4.0
	9.66	$r \leq 2$	15.2
	28.93	$r \leq 1$	29.5
	58.92	$r = 0$	47.2
	Cointegrating Vectors		
FNY	1.000	1.000	
FL	-0.692	-0.972	
SNY	-1.048	-0.216	
SL	0.709	-1.564	
	Statistics of the Error Process		
	Q(6)	Q(12)	JB-Test
FNY	4.28	9.71	0.391
FL	2.87	7.53	0.621
SNY	3.02	9.28	1.442
SL	2.99	13.28	0.328

Note: The statistics of the Likelihood-Ratio test for $\mu = \tilde{\alpha}_0\beta$ and for the dummy follow χ^2-distributions with 1 and 4 degrees of freedom, respectively. The test statistics of the Jarque-Bera (JB) normality test and the Ljung-Box Q(6) and Q(12) for residual autocorrelation are all χ^2-distributed with 2, 6 and 12 degrees of freedom. The 5% critical values for the above χ^2-statistics are 5.99, 9.48, 12.59 and 21.03, respectively.

TABLE 4: ML-COINTEGRATION ANALYSIS OF MODEL (6)

LR-test of $\mu = \tilde{\alpha}_0\beta$ (df = 1)			0.92
LR-test of dummy (df = 4)			27.32
	Eigenvalues		
0.255	0.153	0.109	0.061
	Trace Statistics		Critical Values
	7.31	$r \leq 3$	7.8
	23.63	$r \leq 2$	23.4
	45.57	$r \leq 1$	41.2
	84.46	$r = 0$	63.0
	Cointegrating Vectors		
FNY	1.00	1.00	1.00
FL	-0.52	-1.07	-0.71
SNY	-1.49	-0.27	-0.22
SL	0.93	0.31	-0.11
$r_U S$	0.20	0.00	0.09
$r_U K$	-0.25	-0.07	-0.16
	Statistics of the Error Process		
	Q(6)	Q(12)	JB-Test
FNY	4.08	10.06	0.49
FL	2.57	7.81	0.47
SNY	2.03	5.54	2.99
SL	2.84	12.55	0.16

Note: see Table 3

TABLE 5: ESTIMATES OF COINTEGRATING VECTORS

Based on the Just Identifying Restrictions H_0^β

	Cointegrating Vectors β			Adjustment Coefficients $\tilde{\alpha}$		
	β_S	β_{NY}	β_L	β_S	β_{NY}	β_L
FNY	0.000	1.000	0.000	-0.559	-0.366	0.468
FL	0.000	0.000	1.000	-0.251	-0.071	0.165
SNY	1.000	-1.127	0.000	-0.515	0.116	0.138
SL	-0.983	0.000	-1.082	-0.311	-0.138	0.359
r_{US}	-0.041	0.321	0.268			
r_{US}	-0.032	-0.611	-0.670			
Constant	0.009	0.251	0.188			

TABLE 6: LR-TESTS ON THE STRUCTURE OF THE COINTEGRATING VECTORS

(Given The Identifying Restrictions H_0^β)

	Restrictions	Likelihood Ratio	df
H_1^β	$\beta_{14} = \beta_{23} = \beta_{34} = 1$	7.69 (.06)	3
$H_{r,1}^\beta$	$\beta_{15} = \beta_{16} = 0$	0.46 (.79)	2
$H_{r,23}^\beta$	$\beta_{25} = \beta_{26} = \beta_{35} = \beta_{36} = 0$	8.43 (.07)	4
$H_{c,1}^\beta$	$\beta_{17} = 0$	0.01 (.93)	1
$H_{c2,3}^\beta$	$\beta_{27} = \beta_{37} = 0$	15.19 (.001)	2
H_{rc}^β	$\beta_{15} = \beta_{16} = \beta_{17} = \beta_{26} = \beta_{35} = 0$	6.97 (.22)	5

TABLE 7: WALD TESTS ON ADJUSTMENT COEFFICIENTS

Zero Restrictions on Particular Rows of $\tilde{\alpha}$						
		df	FNY	FL	SNY	SL
H_i^α	$\tilde{\alpha}_{i,j} = 0$	3	3.87 (.27)	1.64 (.64)	12.38 (.01)	7.20 (.06)
$H_{i,1}^\alpha$	$\tilde{\alpha}_{i,1} = 0$	1	2.52 (.11)	0.55 (.45)	2.47 (.11)	0.77 (.76)
$H_{i,23}^\alpha$	$\tilde{\alpha}_{i,2} = \tilde{\alpha}_{i,3} = 0$	2	3.25 (.19)	1.41 (.49)	7.96 (.01)	6.70 (.03)

Note: The tests are conducted using restricted estimates of β according to H_0^β of Table 7.

TABLE 8: GRANGER CAUSALITY TESTS OF COMBINED HYPOTHESES

	Restrictions	Wald-statistics	df
H_S^α	$\tilde{\alpha}_{2,2} = \tilde{\alpha}_{2,3} = \tilde{\alpha}_{3,2} = \tilde{\alpha}_{3,3} = 0$	9.12 (.05)	4
H_S^a	$\tilde{a}_{3,1} = \tilde{a}_{3,2} = \tilde{a}_{4,1} = \tilde{a}_{4,2} = 0$	28.54 (.02)	16
H_F^α	$\tilde{\alpha}_{1,2} = \tilde{\alpha}_{1,3} = \tilde{\alpha}_{2,2} = \tilde{\alpha}_{2,3} = 0$	1.32 (.86)	4
H_F^a	$\tilde{a}_{1,1} = \tilde{a}_{1,2} = \tilde{a}_{2,1} = \tilde{a}_{2,2} = 0$	22.94 (.12)	16
H_{NY}^α	$\tilde{\alpha}_{1,1} = \tilde{\alpha}_{1,3} = \tilde{\alpha}_{3,1} = \tilde{\alpha}_{3,3} = 0$	30.06 (.001)	4
H_{NY}^a	$\tilde{a}_{1,2} = \tilde{a}_{1,4} = \tilde{a}_{3,2} = \tilde{a}_{3,4} = 0$	47.14 (.001)	16
H_L^α	$\tilde{\alpha}_{2,1} = \tilde{\alpha}_{2,2} = \tilde{\alpha}_{4,1} = \tilde{\alpha}_{4,2} = 0$	1.52 (.82)	4
H_L^a	$\tilde{a}_{2,1} = \tilde{a}_{2,2} = \tilde{a}_{4,1} = \tilde{a}_{4,2} = 0$	30.81 (.01)	16
$H_{NY;F}^\alpha$	$\tilde{\alpha}_{1,2} = \tilde{\alpha}_{1,3} = \tilde{\alpha}_{2,1} = \tilde{\alpha}_{2,2} =$ $\tilde{\alpha}_{2,3} = \tilde{\alpha}_{3,2} = \tilde{\alpha}_{3,3} = 0$	2.40 (.93)	7
$H_{NY;F}^a$	$\tilde{a}_{1,3} = \tilde{a}_{1,4} = \tilde{a}_{2,1} = \tilde{a}_{2,3} =$ $\tilde{a}_{2,4} = \tilde{a}_{4,1} = \tilde{a}_{4,3} = 0$	51.12 (0.004)	28

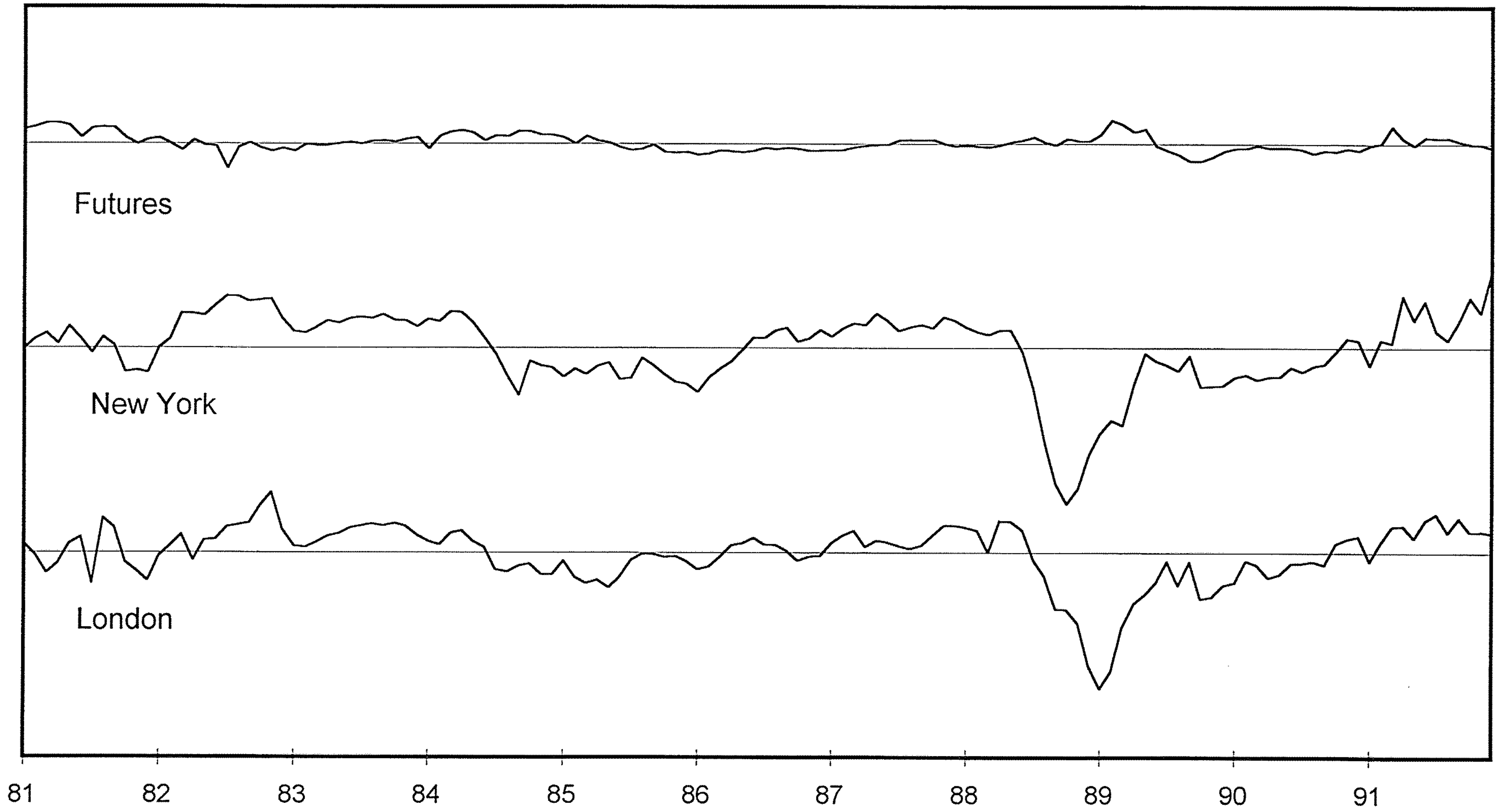
Note: The tests are conducted using restricted estimates of β according to H_0^β of Table 7. The tests for H_F^α and $H_{F;NY}^\alpha$ are based on a different representation of β (see text). $\tilde{a}_{ij}$ refers to the i-th row and j-th column of the lag polynomial $\tilde{A}_{11}^*(L)$ in equation (7). The Wald-statistics for $H_{(.)}^a$ follow a χ^2-distribution with 16df (for $H_{NY;F}^a$ 28 df) under the null.

TABLE 9: RESTRICTED ESTIMATES OF COINTEGRATING VECTORS

Based on Restrictions H^{β}_{rc} and $H^{\alpha}_{NY;F}$

	Cointegrating Vectors β			Adjustment Coefficients $\tilde{\alpha}$		
	β_F	β_{NY}	β_L	β_F	β_{NY}	β_L
FNY	1.000	1.000	0.000	-0.217	0.000	0.000
FL	-1.000	0.000	1.000	0.000	0.000	0.000
SNY	0.011	-1.011	0.000	-0.152	0.492	-0.320
SL	0.036	0.000	-0.956	0.000	0.000	0.145
r_{US}	0.063	0.063	0.000			
r_{UK}	-0.057	0.000	0.057			
Constant	0.091	0.033	-0.059			

Figure 1: Cointegration Relationships



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